INTERRUPTED WORK CAREERS: DEPRECIATION AND RESTORATION OF HUMAN CAPITAL*

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ABSTRACT

The quantitative effects and even the existence of a "human capital depreciation" phenomenon have been a subject of controversy in the recent literature. Prior work, however, was largely cross-sectional and the longitudinal dimension, if any, was retrospective. Using longitudinal panel data (on married women in NLS) we have now established that real wages at reentry are, indeed, lower than at the point of labor force withdrawal, and the decline in wages is greater, the longer the interruption. Another striking finding is a relatively rapid growth in wages after the return to work. This rapid growth appears to reflect the restoration (or "repair") of previously eroded human capital. The phenomenon of "depreciation" and "restoration" is also visible in data for immigrants to the United States. However, while immigrants eventually catch up with and often surpass natives, returnees from the nonmarket do not fully restore their earnings potential.

I. INTRODUCTION

Interruption of a work career, especially a lengthy one, can be expected to reduce a person's earning power. The quantitative effects and even the existence of this "human capital depreciation" phenomenon have been a subject of controversy in the recent literature.[1] Prior work, however, was

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1 See especially Mincer and Polachek [7, 8], Sandell and Shapiro [9], and Corcoran and Duncan [5].

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largely cross-sectional, and the longitudinal dimension, if any, was limited to retrospectively reported duration of past interruptions.

Using National Longitudinal Survey (NLS) panel data on wages of married women,[2] we are now able to establish the existence of depreciation phenomena more firmly. We observe that real wages at reentry are on the average lower than at the point of labor market withdrawal, and the decline is greater, the longer the interruption. If wage setbacks due to interruptions were attributable solely to forgone growth of transferable ("general") human capital, wages of returnees would be lower than wages of stayers, but not lower than their own wages at exit. Wages at reentry would be lower than at exit to the extent that capital specific to the last job was lost by separating. But if this only were the case, the decline in wages would not depend on the length of the interruption.[3]

Although interruption periods typically mark sharp declines in earning power, there is also a relatively rapid initial growth in wages after the return to work. It is rather surprising to find that returnees from the nonmarket appear to incur greater job investments upon return to the market than do stayers of the same age and education levels who just changed jobs, although the returnees may invest more after than before the interruption.

Our interpretation of this wage "rebound" phenomenon is restoration or "repair" of the previously eroded human capital, on the assumption that reconstruction of occupational skills is more efficient—that is, less costly—than new construction of human capital. In effect, the rate of decline in the rental value of the depreciated stock is greater than its rate of "physical" depreciation. This is because the market productivity of the eroded stock is greatly reduced even if only small parts become defective. It may take relatively little effort to replace or to repair the defective parts, and once accomplished the rental value rises to its normal rate.

Put in terms of the Ben-Porath model of production of human capital [2], the erosion of human capital reduces its market productivity more than its productivity as an input in its own reproduction. Consequently, the opportunity costs of investments of leavers are smaller, upon reentry, than are those of comparable stayers. The lower costs induce initially larger

2 We did not explore work interruptions of men in the present study. They tend to be infrequent and quite short. The longer ones are usually a matter of health, schooling, or the military. See Corcoran and Duncan [5] for a more detailed discussion of sex differences in work interruptions.

3 See the discussion in Section V below. Our evidence contradicts the conclusions of the Corcoran-Duncan study to the effect that "labor force withdrawals do reduce wages because work experience is not being accumulated, but there is no *additional* penalty due to depreciation of skills" [5, p. 18]. Their study utilizes a different data set (PSID) and does not exploit longitudinal information on wages.

investments and steeper wage growth. But after the earning capacity of returnees is restored to the preinterruption level (point F in Figure 1), their investment profile becomes identical with, hence earnings parallel to, the profile of stayers. The postrestoration profile may, of course, be flatter if further interruptions are expected.

The fact that wages grow rapidly upon return to work suggests that different estimates of depreciation rates can be obtained depending on the period of observation. The estimate is smaller if wages of returnees are observed years after the interruption spell than immediately after it. But the long-run depreciation effect is never zero as long as the recovery is not instantaneous ($F \neq E$ in Figure 1). Long-run depreciation is larger, the longer the recovery period. The distinction between the short and long run may account for some of the variation in findings reported in the literature.

The phenomenon of depreciation and restoration of human capital is not restricted to intermittent workers. An interesting example is the economic experience of international migrants. Recent studies of wages of immigrants to the U.S. reveal comparable patterns of decline and increase in occupational status before and after their arrival in the U.S.[4] We review some of these findings in the light of our hypothesis in the last section of the paper.

II. THE INTERRUPTED EARNING PROFILE: A WORKING SCHEME

A highly simplified graphic representation of the analysis that follows is shown in Figure 1. Age-earning profiles for a continuous worker and an intermittent worker are given by the straight line JKL and by the kinked line $ABCDEFG$, respectively. For simplicity it is assumed that the intermittent worker experiences only one interruption which lasts only one period of time.[5] We may thus distinguish four principal phases in the work and wage history of such a worker: a preinterruption period (AB), the interruption ($BCDE$), a restoration period (EF), and a postrestoration period (FG).

More generally, interruptions may be repetitive or unique, unexpected or planned, or at least vaguely anticipated. A spectrum of such patterns can be found in work histories of married women (Mincer and Polachek

4 See Chiswick [3, 4] and De Freitas [6].

5 The assumption of "one period" interruption ($T - V = 1$) is an expository convenience soon to be relaxed. In the meantime it serves to express the slopes of the age-earning profile (and thus the coefficients of the earning function) in terms of vertical intervals in Figure 1, but otherwise does not affect the fundamentals of the analysis or its generality. Linearity of the profile is also a matter of convenience.

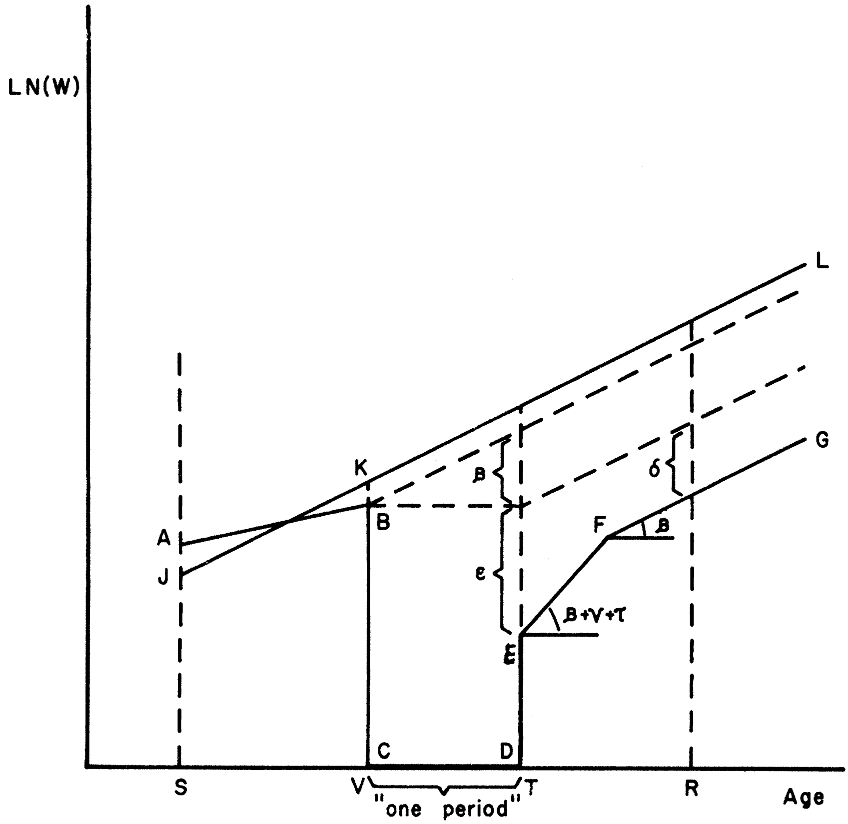

FIGURE 1

Note: If the interruption is taken to be "one period" long (e.g., $T - V = 1$), then the vertical interval β is equivalent to the rate of depreciation due to lost experience; ϵ is the rate of depreciation due to nonuse (holding experience constant) in the *short run*; and δ is the rate of depreciation due to nonuse in the *long run.* Note that ϵ is reduced in the long run to δ by the combined effects of two countervailing processes typical to the postinterruption period: restoration of market productivity γ, and accumulation of job tenure τ.

[7]). If the interruption is anticipated, the preinterruption period may show a relatively flat wage profile (segment AB rather than JK in Figure 1) which reflects lower rates of investment in human capital. Work prior to interruptions occasioned by planned family formation is, or used to be, a typical example.

In the presentation of the second phase, the interruption itself, we show wages at the reentry point (DE) to be below wages at the point of

TABLE 1
WITHDRAWAL AND REENTRY WAGE RATES BY THE LENGTH OF THE INTERRUPTION PERIOD (1967 PRICES) NLS PANEL OF MARRIED WOMEN AGES 30–44 IN 1967

Years of Nonparticipation	0	1–2	3–4	5–6
Withdrawal hourly wage rates ($)	2.27[a]	1.92	1.70	1.73
Reentry hourly wage rates ($)	2.35[a]	1.75	1.46	1.27
Number of observations	931	128	141	104

a $2.27 and $2.35 are average hourly wage rates for continuous workers in 1971 and 1972 (1967 prices), respectively.

labor force withdrawal (*CB*). This finding can already be seen from a preliminary inspection of the NLS data. Table 1 reports average hourly wage rates immediately before the interruption, and immediately after reentry, for groups consisting of the *same individuals* at the two points. (Obviously, such comparisons of wages at different points in time are possible only with longitudinal data.) It is evident that reentry wages fall short of withdrawal wages, and that the gap increases with the duration of the interruption. This further indicates that the lower wages earned by intermittent workers are not only a result of lost experience during the interruption and less investment during the preinterruption periods, but also as a result of deterioration of earning power due to nonuse.

Another noteworthy feature of Table 1 is the lower preinterruption wage of workers who interrupt for longer periods. If intermittency is anticipated or repetitive, the lower preinterruption wage is a consequence of lesser investment in human capital. At the same time, an upward slope in the lifetime supply of labor predicts that lower wage workers will interrupt their work careers for longer periods and more frequently. In the cross-section data where the distinction between prior and subsequent wages is not available, it can be argued that the negative relation between interruptions and wages is really a supply effect of (prior) wages rather than the effect of interruption on (subsequent) wages. The longitudinal data clearly establish the latter effects in addition to and as distinct from labor supply effects.

The period immediately after the interruption, the third phase, is associated with a rapid process of wage growth. As we infer once again from a preliminary inspection of data, this time in Table 2, a rapid restoration process of earning power takes place (roughly) throughout the first five years after reentry. We further note in Table 2 that the rapid wage growth during that phase is associated with accumulation of job tenure. (It

TABLE 2
HOURLY WAGE RATES OF INTERMITTENT WORKERS BY CURRENT EXPERIENCE AND BY JOB TENURE

Years Since Last Interruption (1)	Job Tenure (in Years) (2)	Hourly Wage Rate (1967 Prices) (3)	Number of Observations (4)
1	.9	\$2.61	101
2	1.2	2.75	238
3	1.7	2.78	156
4	2.8	2.94	148
5	2.9	3.10	209
6	3.6	3.07	120
7	4.6	3.23	83
8	4.5	3.16	78

appears, however, from our subsequent analysis that the tenure-related wage growth accounts for the lesser part of the observed wage progress.) The growth in wages of returnees should eventually level off and settle at a rate similar to that of continuous workers, or lower if further interruptions are anticipated.[6] This point marks the beginning of the fourth and last phase of the process outlined above.

Some aspects of the interrupted earning profile can be parameterized and estimated. In what follows, the short-run and long-run effects of nonparticipation (ϵ and δ in Figure 1) and the long-run effect of experience (β) are first estimated on the basis of retrospective data and then reestimated along with the short-run effect of experience (γ) and tenure (τ) from panel data.

III. ESTIMATIONS FROM RETROSPECTIVE DATA

All the working estimators in the following analysis are essentially special cases of the earning function

$$\ln(w) = \alpha s + \beta e_0 + \gamma e_1 + \delta h_0 + \epsilon h_1 + \mu x \tag{1}$$

where the logarithm of wages $\ln(w)$ is specified as a function of two pairs of

6 The parallelism of the last phase (of *FG* to *KL* in Figure 1) should, strictly speaking, hold for dollar profiles, if beyond *F* intermittent workers invest the same amounts as continuous workers. In Figure 1 wages are drawn in logarithms, so the parallelism denotes somewhat lesser investments by intermittent workers.

experience-nonexperience variables: past and current labor force participation (e_0 and e_1, respectively); past and current nonparticipation (h_0 and h_1, respectively).[7] We may thus interpret the coefficients β and γ as the long-run and short-run effects of participation, and δ and ϵ as the long-run and short-run effects of nonparticipation. All these effects are controlled in equation (1) for schooling s and for a vector x of "other" variables to be discussed later on.

Wages, of course, are not observable for individuals currently out of the labor force, except at the point of reentry when current experience is still almost zero. Estimation at reentry points is possible with longitudinal data because individuals observed at their labor force withdrawal and reentry points are identifiable.

Let the subscript T denote the timing of the most recent labor force reentry (termination of the interruption period) for an intermittent worker. Then, equation (1) becomes

$$\ln(w_T) = \alpha s + \beta e_0 + \gamma h_0 + \epsilon h_1 + \mu x_T \tag{2}$$

because $e_1 = 0$. This particular specification enables us to estimate the long-run effects of prior experience (e_0) and of prior interruptions (h_0) as well as the short-run effect of the just completed interruption (h_1). Note, however, that the reentry point occurs at different calendar times for different individuals, and thus the observations must be aligned (according to T) and appropriately adjusted for chronological differences such as inflation and age. Wages are adjusted for inflation using 1967 prices. Age and all the experience variables are chronologically adjusted by the identity Age $= 6 + s + e_0 + h_0 + h$, and thus controlled in equation (2).

The Long-Run Effects

Equation (2) has been fitted to the NLS (mature women 1966–1974) data, and the results are reported in Table 3. There are two major findings. First, the long-run effects of market experience and nonexperience are both statistically significant in the predicted direction. These further indicate not only that experience and nonexperience have lagged effects on wages, but also that the effects persist throughout and last beyond spells of labor force withdrawal, which typically involve new jobs and new employers.

Quantitatively, we find in the *long run* between .6 and 1.1 percent

7 The precise definitions used in the construction of these variables are: e_0 = years of experience prior to the most recent spell of nonparticipation; e_1 = years of participation after the most recent spell of nonparticipation; h_0 = years of nonparticipation prior to the most recent spell of employment; and h_1 = years of nonparticipation after the most recent spell of employment. So defined, strictly positive values for both e_1 and h_1 will never occur in the same observation.

TABLE 3
ESTIMATES OF EQUATION 2 (NLS DATA[a])

Sample[b]	I (1)	I (2)	II (3)	II (4)
s	.057	.052	.037	.037
	(8.49)	(7.85)	(2.74)	(2.74)
h_1	−.076	−.060	−.033	−.045
	(9.23)	(6.95)	(1.97)	(2.39)
h_0	−.006	−.006	−.011	−.010
	(2.08)	(1.90)	(1.87)	(1.75)
e_0	.018	.018	.020	.022
	(2.47)	(2.42)	(1.49)	(1.62)
e_0^2	−.00027	−.00038	−.00096	−.00096
	(1.14)	(1.59)	(1.71)	(1.72)
TEN	—	.016	—	.008
		(5.21)		(.33)
LYOF	—	−.258	—	−.274
		(5.70)		(3.50)
UNEM	—	.069	—	.149
		(1.46)		(1.71)
MAR	—	−.048	—	−.135
		(.63)		(.80)
DIV	—	.033	—	.001
		(.64)		(.00)
BAB	—	.033	—	−.000
		(.77)		(.00)
HLTH	—	−.072	—	−.107
		(2.10)		(1.70)
MIG	—	−.032	—	−.091
		(.92)		(1.45)
Const.	4.52	4.56	4.69	4.79
R^2	.19	.24	.05	.08
N	1485	1485	612	612

a For definitions of variables, means, and standard deviations, see the glossary and the summary of statistics in the appendix.
b For a description of samples, see footnote to Appendix Table 1.

decline in wages per year of nonparticipation (h_0), depending on whether the estimates are derived from the sample which includes women who did not interrupt their careers (sample I, cols. 1 and 2) or the sample is confined to women who did (sample II, cols. 3 and 4).[8] Similarly, a year of experience

8 For a description of the various samples used, see footnote to Appendix Table 1.

(e_0) results in a long-run increase in wages (calculated at the mean) ranging between .4 percent (sample II, cols. 3 and 4) and 1.2 percent (sample I, cols. 1 and 2). It should be emphasized that all these estimates represent partial effects: the long-run effect of nonparticipation is estimated holding experience constant, and the long-run effect of experience is estimated holding nonparticipation constant. It follows that in order to evaluate the total effect of work interruption the two effects should be summed. Thus, the total cost of a year outside the labor force ranges between 1.5 percent (= 1.1 + .4, for sample II) and 1.8 percent (= .6 + 1.2, for sample I) of wage decline in the long run.

The Short-Run Effect of Nonparticipation

The second major finding is that the cost of nonparticipation is substantially higher in the short run. The short-run effect on wages of current nonparticipation (h_1) is estimated to range between 3.3 and 7.6 percent per year depending on the specification and the sample used. These figures overstate the loss of general human capital since they include the effect of forgone tenure. The latter raises wages of the combined sample by 1.6 percent at the mean (col. 2, Table 3) and of the returnees by 1.1 percent (judging by their postinterruption experience in col. 3, Table 7). Thus, growth of tenure accounts for a small part of the short-run wage rebound. The latter is by definition equal, but opposite in sign, to the estimated short-run depreciation effect, while the former corresponds to the loss of specific capital (at mean tenure levels). If so, the loss of specific capital accounts for at most one-third of total depreciation.

The Nature and Length of the Interruption Spell

We note that interruptions associated with layoff (*LYOF*), ill health (*HLTH*), and migration (*MIG*) result in greater than average depreciation. The nature of these differences is not explored further in this paper. These and other events which often cause withdrawals of women from the labor market also appear to affect the *duration* of the interruption spell, as shown in the regression results reported in Table 4. Included in this estimation are all the NLS respondents who have experienced a complete interruption spell within the survey period (1966–1974). While the average interruption is 2.7 years long (with a standard deviation of 1.6 years), the findings indicate that child bearing (*BAB*) is associated with significantly longer interruptions. Shorter than the average are the interruptions associated with divorce (*DIV*), layoff (*LYOF*) and unemployment (*UNEM*). The duration of interruptions associated with the occurrence of marriage (*MAR*), ill health (*HLTH*), and migration (*MIG*) seem to differ little from the average.

TABLE 4
ESTIMATED EFFECTS ON THE DURATION OF THE INTERRUPTION PERIOD[a]

	b (1)	t (2)
s	−.045	(1.57)
DIV	−.665	(2.65)
MAR	−.075	(.19)
BAB	.707	(4.28)
HLTH	.145	(.95)
MIG	.061	(.41)
UNEM	−.306	(1.53)
LYOF	−.866	(4.59)
Const.	3.51	
N	706	
R^2	.22	

a For glossary of variables and summary of statistics (means and standard deviations), see the Statistical Appendix.

We further note that the duration of the interruption is inversely related to the level of education. This is consistent with a positively sloping lifetime labor supply function. It is also consistent with previous findings that the depreciation rate increases with the level of education (Mincer and Polachek [7], Tables 5 and 6; also Mincer and Polachek [8], Tables 1 and 2),[9] thereby deterring the more educated from lengthier episodes of nonparticipation.

IV. ESTIMATION FROM PANEL DATA

Reestimating the Short Run Effects

So far, the effects of nonparticipation have been estimated partly on the basis of retrospective data.[10] In this section we dispense with retrospective data. To this end we replace the retrospective experience-nonexperience variables (e_0, h_0) with information on preinterruption wages available in

9 Corcoran and Duncan [5] did not find any relation between the depreciation rates and occupation in the Panel Study of Income Dynamics. They do not report comparisons by education level.

10 Since the NLS panel started in 1967, all prior data are retrospective. Thus e_0 and h_0 are largely retrospective.

TABLE 5
ESTIMATES OF EQUATION 4 (NLS DATA[a])

Sample[b]	III (1)	III (2)	IV (3)	IV (4)
s	.028	.028	.009	.008
	(6.47)	(6.47)	(.869)	(.806)
h_1	−.089	−.086	−.068	−.072
	(15.3)	(14.3)	(4.94)	(4.94)
$\ln(w_0)$	.585	.578	.684	.673
	(34.0)	(33.1)	(23.2)	(22.6)
LYOF	—	−.071	—	−.101
		(2.04)	—	(1.49)
UNEM	—	.045	—	.134
		(1.30)		(1.84)
MAR	—	−.011	—	−.054
		(.232)		(.500)
DIV	—	.012	—	.015
		(.320)		(.145)
BAB	—	−.019	—	−.005
		(.619)		(.089)
HLTH	—	−.029	—	−.058
		(1.16)		(1.10)
MIG	—	−.061	—	−.087
		(2.36)		(1.67)
Const.	1.91	1.97	1.56	1.67
R^2	.60	.61	.62	.63
N	1304	1304	373	373

a Means, standard deviations, and definitions for all the variables are given by the glossary and summary of statistics, in the Statistical Appendix.
b For a description of samples, see footnote to Appendix Table 1, Statistical Appendix.

the panel data. Here the specification of the earning function involves the following steps:

Let the subscripts V and T denote years of labor force withdrawal and reentry associated with the most recent interruption. Evaluated at the point V, the original earning function (1) becomes:

$$\ln(w_V) = \alpha s + \beta e_0 + \delta h_0 + \mu x_V \tag{3}$$

since both $e_1 = 0$ and $h_1 = 0$. Substituting (3) into (2) we obtain

$$\ln(w_T) = \lambda \ln(w_V) + \epsilon h_1 + \mu(x_T - x_V) \tag{4}$$

where $\lambda = 1$. Alternatively, we may subtract (3) from (2) to obtain

$$\ln(w_T) - \ln(w_V) = \epsilon h_1 + \mu(x_T - x_V) \tag{5}$$

TABLE 6
ESTIMATES OF EQUATION 5 (NLS DATA[a])

Sample[b]	III (1)	III (2)	IV (3)	IV (4)
s	.004	.006	.008	.008
	(.85)	(1.19)	(.75)	(.659)
h_1	−.059	−.056	−.064	−.060
	(8.60)	(7.81)	(4.04)	(3.60)
LYOF	—	.074	—	.019
		(1.79)		(.251)
UNEM	—	.041	—	.082
		(.976)		(.980)
MAR	—	−.038	—	−.063
		(.64)		(.501)
DIV	—	−.004	—	.009
		(.08)		(.077)
BAB	—	−.050	—	−.030
		(1.33)		(.46)
HLTH	—	.033	—	.041
		(.10)		(.673)
MIG	—	−.051	—	−.071
		(1.62)		(1.19)
Const.	−.009	−.028	.159	.162
R^2	.06	.06	.04	.05
N	1304	1304	373	373

a For means, standard deviations, and definitions of variables, see the glossary and summary of statistics (Appendix Table 1), the Statistical Appendix.
b For description of samples, see footnote to Appendix Table 1, Statistical Appendix.

Deterministically, (4) and (5) are equivalent. Stochastically, they differ because of differences in the error term. Both specifications permit estimation of the short-run effect of interruption, ϵ, without a need to resort to retrospective data. Instead, they make more efficient use of the current data: wages at the most recent points of withdrawal and reentry (w_V and w_T), the duration of the most recent interruption (h_1), and recent changes in the "other" variable ($x_T - x_V$). Panel estimates of equations (4) and (5) are reported in Tables 5 and 6, respectively. The estimates of ϵ are similar in both tables and range between 5.9 and 8.9 percent. These results are comparable and a bit higher than the 3.3–7.6 range obtained in Table 4 on the basis of retrospective data. Thus they confirm and reinforce the findings of the previous section, namely, that the size of the short-run effects of nonparticipation greatly exceed the long-run effects and are larger than any reported in previous studies.

It is worth noting that the estimate of λ (the coefficient on $\ln(w_V)$), which was not constrained to equal 1, is much less than 1. Unless errors of measurement in wages are large, the estimate suggests larger (percent) losses in wages at higher wage levels, given the length of interruption.

Experience and Job Tenure in the Postinterruption Period

In order to analyze the process of wage growth in the postinterruption period, we now evaluate the earnings function (1) at a fixed chronological date: the last year of the panel (1974) in our sample. Here (1) becomes

$$\ln(w_{74}) = \alpha s + \beta e + \gamma e_1 + \delta h_0 + \tau(TEN) + \mu x \tag{6}$$

where $h_1 = 0$ and $e = e_0 + e_1$. Information on tenure (TEN), available for the postinterruption period, was explicitly included in the specification of (6).[11] Wage growth during the postinterruption period is the sum of the three coefficients $\beta + \gamma + \tau$ (equal to the short-run effect of current experience). Equation (6) has been fitted to the data (samples V and VI) and the results are summarized in Table 7. Based on the results estimated by the linear form (cols. 1, 3), postinterruption wages tend to grow at an average rate of roughly 2.5 percent per year of experience. The quadratic form (cols. 2, 4) results are somewhat higher: 3.3 to 3.6 percent at the mean (which occurs about nine years after the most recent interrruption for the average respondent).

When these estimates are projected down to the first year after reentry, the rate of growth, as expected, is much higher: between 5.8 and 6.4 percent per year. This rate of growth of wages of women who were past the age of 30 at reentry is almost double the estimated rate for men (3.4 percent) of the same age, projected to the outset of their working lives (see Mincer and Polachek [7], Table 11). It seems more reasonable to us to view this large difference in wage growth as a difference in the nature rather than in the scale of the human capital investment. A breakdown by factors indicates that in all the above cases job tenure accounts for less than half of the postinterruption wage growth. More than half of it can be ascribed to experience net of the effect of tenure: namely, to growth of earning power embodied in the worker (general training) and transferable with him across jobs. In turn, a half or more of the latter is due to the repair of human capital (coefficient of e_1 holding e constant).

In addition to the various short-run effects outlined above, equation

11 Tenure may be longer than the postinterruption period for workers who returned to the same employer.

TABLE 7
ESTIMATES OF EQUATION 6 (NLS DATA[a])

Sample[b]	V (1)	V (2)	VI (3)	VI (4)
s	.057	.056	.064	.062
	(9.38)	(9.27)	(9.18)	(8.99)
h_0	−.0083	−.024	−.0039	−.017
	(2.88)	(3.86)	(1.22)	(2.22)
h_0^2	—	.0007	—	.0005
		(2.90)		(1.82)
e	.0042	.019	.0073	.015
	(1.43)	(2.00)	(2.19)	(1.45)
e^2	—	−.0004	—	−.0002
		(1.77)		(.31)
e_1	.0093	.023	.0084	.024
	(2.83)	(2.83)	(2.26)	(2.70)
e_1^2	—	−.0005	—	−.0006
		(1.56)		(1.69)
TEN	.012	.022	.011	.019
	(4.11)	(3.06)	(3.41)	(2.32)
TEN^2	—	−.0005	—	−.0003
		(1.40)		(.811)
Const.	4.88	4.73	4.70	4.63
R^2	.28	.29	.25	.27
N	1015	1015	820	820

a For means, standard deviations, and definitions, see the glossary and summary of statistics (Appendix Table 1), Statistical Appendix.
b For a description of samples, see footnote to Appendix Table 1, Statistical Appendix.

(6) offers reestimation of the long-run effects of experience (e) and nonexperience (h_0). The findings of .4 to .8 for the former, and −.4 to −1.0 for the latter, are in clear agreement with the findings obtained in the previous specifications (Section III).

V. *SOME CONCLUSIONS, CONJECTURES, AND THE CASE OF IMMIGRANTS*

Our short- and long-run depreciation rates are linear estimates representing the average loss of earning power due to an additional year of nonwork. In longitudinal data the negative coefficient of h_1 is, indeed, evidence on the existence of depreciation. We would expect that depreciation affects both general and specific human capital. But, while losses of general capital increase with the duration of absence from work, the loss of specific capital is a once-for-all phenomenon due to separation from the job. This means that, if the losses were only in specific capital, the correctly estimated marginal depreciation rates would be zero. Consequently, we may reject the notion that observed depreciation rates are restricted to specific capital. In previous work (Table 3 in [8]), we found that interruptions not exceeding a year had negligible, insignificant effects on wages. Apparently intermittent workers lose little in specific capital, probably because they accumulate little of it. Since we find that wages decline as a function of duration of the interruption, what we are observing is largely the phenomenon of depreciation and restoration of general human capital.[12]

Although we estimated only linear effects of interruptions, we should not conclude that the depreciation *rate* is independent of the duration of the interruption. Longer periods of absence may well accelerate losses of knowledge and skill. Beyond some point, with a substantial part of the stock gone, additional losses may well diminish. Of course, when all the skill has been forgotten and lost, no further erosion is possible. We may expect that, if observable over a long range of interruption periods, the depreciation rate would be a growth (decay) type function ($\int$) of duration. In the previously referred work [8], we estimated separate (dummy) coefficients for interruptions of one, two, and three plus years, and these showed negative coefficients increasing in size. This is consistent with the initial (accelerating) part of the proposed $\int$ curve.

Empirical estimation of the complete pattern of depreciation is difficult for several reasons. Much longer panel data would be required to accommodate progressively lengthy interruptions. But even with very long panels, it is unlikely that very long interruptions would be observed. This is because the longer the stay in the nonmarket, the less likely the return to the market. Although the interruption may result from an increase in the shadow wage (e.g., increased family demand) above the market wage, a long decline in the latter (due to depreciation) may well leave it below the shadow wage even when the shadow wage returns to its usual level. At the

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12 Is it not possible that the larger depreciation observed in longer withdrawals results merely from the loss of specific capital, given the not unreasonable assumption that those with a shorter interruption are more likely to return to the prior employer, thereby not losing any specific capital? This assumption is, indeed, valid. However, as we already estimated (the second subsection in Section III above), the maximal loss of specific capital is much less than the total loss. Moreover, the number of women who returned to the same employer after two years of interruption is nil in our sample. Yet, Table 2 shows a continuous and strong decline in wages between three and six years of interruption.

same time nonmarket skills may increase with nonmarket experience, thereby raising the "normal" level of the reservation wage.

Moreover, the longer the interruption that is observed, the more likely it is that the returnees are people whose human capital is especially durable, whether it is a matter of personal resiliency (good memory) or environment (lesser changes in the field, or special opportunities that have arisen for them). Consequently, our estimates of short- and long-run depreciation, observed on returnees (prior to 1974) only, are likely to be understated. By the same token, the subsequent wage growth is probably overstated. Nevertheless, the ultimately lower wage level of returnees compared to stayers indicates that to a sufficiently large extent the wage rebound after interrruption differs qualitatively from the usual continuous investment trajectory.

Partial losses of human capital may result from causes other than interruption of market work by nonmarket activities. An interesting example is international migration. Skills and knowledge are not completely transferable across frontiers. The greater the economic and cultural "distance" between country of origin and of destination, the greater the "depreciation" of human capital. Here "distance" plays the same role as duration of work interruption in the case of intermittent workers. Since not all skills are equally affected, economically motivated migrants are likely to be selected from among the most adaptable persons and those with the most adaptable skills. Selectivity by occupational skill, though not by personal motivation and stamina, is likely to be weaker in extra-economic migration, as in the case of political or religious refugees. Therefore, greater losses of human capital may be experienced by them. This is analogous to the experience of intermittent workers: returnees to the market are also likely to be those who lost least by interrupting, and greater losses, on average, can be expected when the interruption is unanticipated.

Just as in the case of returnees to the labor market, new immigrants initially experience the greatest loss in human capital. This is visible in occupational data of immigrants to the U.S. Recent studies (Chiswick [3, 4], De Freitas [6]) have emphasized the strong upward economic mobility of migrants in the U.S. labor market. However, the success story in the U.S. follows an initial drop from the immediately preceding occupational position in the country of origin. According to the 1970 Census data, 22.6 percent of the men arriving in the U.S. between 1965 and 1970 (and in the labor force in 1970) experienced an initial occupational decline, as measured by major occupational categories. As expected on the basis of "distance," the extent of decline (proportion experiencing downward mobility) was 11 percent for immigrants from English-speaking countries, 20.5 percent from other developed countries, and 25.4 percent from LDCs. And, according to Chiswick, the initial decline was largest for immigrants who were predominantly refugees.

According to our interpretation of the behavior of returnees from the nonmarket, readaptation ("repair") of skills is likely to be more efficient than new investments in human capital. The strong upward occupational mobility of immigrants and the steep wage increases during the first half-dozen years in the U.S. partially represent, in our view, the same "rebound" from the decline occasioned by migration as by nonparticipation. Net of the standardizing variables such as education, age, and others, coefficients of years of work experience in the U.S. in immigrant earnings functions exceed the comparable coefficients of U.S. natives, especially at the start of U.S. experience.

We would expect that intermittent workers are lesser lifetime investors (in human capital) than continuous workers. Their wage profiles are indeed lower and flatter than the profiles of continuous workers, despite the temporary steep growth of wages following an interruption. In contrast, economic migrants may well be persons with greater capacity or opportunity than comparable natives at both origin and destination, and greater investors in their human capital.[13]

It is tempting, therefore, to interpret the steeper growth of wages of immigrants than of comparable natives as evidence of their larger investments in human capital [Chiswick, 3, 4]. It is our view, however, that the initially lower wages of immigrants compared to natives and the following initial rapid growth are in part a reflection of "depreciation and restoration" in the rental price of the immigrant's human capital as well as of the larger scale of investment in its stock which is evident as a lifetime proposition. But even with larger investments than those of natives, immigrants from the most "distant" countries, because they suffer the greatest initial differential (e.g., some Asians, and others from LDCs) do not overtake natives in wages, as do migrants from the more industrialized and culturally closer countries (see Chiswick [3, 4]).

Can we distinguish short-term "rebound" from long-term scale of investment? For one, the former implies greater concavity of earnings than does the latter. Second, strength of the rebound should be less sensitive to age which otherwise sharply reduces the volume of investment. Evidence

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13 If greater ability raises both opportunity costs of and the returns to migration, say by an equal proportion, in the presence of direct costs of migration which are independent of ability, migration must be more profitable for the more able workers. (The argument is adopted from Becker [1, p. 76].) Even if abilities did not differ, migration is likely to be selective of people for whom special opportunities beckon at the destination.

TABLE 8
EARNINGS FUNCTIONS OF NATIVE AND FOREIGN-BORN WHITE MEN (ln WAGE, 1969)

	All Native Men		Native Men Who Entered the Labor Force 1960–1969		Foreign-Born Men	
	b	*t*	*b*	*t*	*b*	*t*
s_a	.0689	50.1	.0885	27.6	.0596	5.3
e_a	.0361	32.0	.0910	9.1	.1356	3.2
e_a^2	−.0006	25.2	−.0032	3.5	−.0080	1.9
s_f					.0501	21.8
e_f					.0200	5.7
e_f^2					−.0003	4.5
$e_a \cdot e_f$					−.0011	3.3
Rural	−.1878	22.1	−.1100	6.3	−.0326	.9
South	−.1313	15.3	−.1190	7.0	−.2113	8.5
Single	−.2252	20.7	−.1645	9.2	−.2080	8.9
R^2	.162		.187		.138	
N	32,933		7,629		5,760	

Source: 1970 Census of Population, as shown in De Freitas [6].
s_a = Years of schooling in the U.S.
s_f = Years of schooling abroad.
e_a = Years of U.S. labor market experience.
e_f = Years of foreign labor market experience.

in favor of the existence of the "restoration" phenomenon may be inferred from Table 8. It provides a comparison of 1970 earnings functions of immigrants to the U.S. with those of native men who entered the labor force between 1960 and 1969. The linear coefficient on the U.S. labor market experience of immigrants is 50 percent larger than the corresponding coefficients of native entrants into the labor force, despite the fact that, on arrival, immigrants were, on average, in their second decade of working life. The deceleration of wages (coefficient of the quadratic term) is more than twice as strong for immigrants. Moreover, the initial growth of wages after arrival in the U.S. is only slightly smaller for older than for younger immigrants; it declines only .1 percent per year (coefficient of interaction term) rather than .8 percent per year as would be indicated by the quadratic term of their U.S. earnings profile. At this pace of decline, initial U.S. wages of 50–60-year-old immigrants still grow more rapidly than wages of U.S. young labor force entrants. It seems apparent that efficient readaptation of the previously acquired human capital stock is an important part of the immigrant success story, as it also seems to be a condition of the reabsorption into the labor market of returness from the nonmarket.

STATISTICAL APPENDIX: GLOSSARY OF VARIABLES

Education

s: Years of schooling.

Wages

w_0: Hourly wage rates at (or immediately before) the most recent labor force withdrawal.

w_1: Hourly wage rates at (or immediately after) the most recent labor reentry.

w_{74}: Hourly wage rates in 1974.

$\Delta\ln(w)$: $\ln(w_1) - \ln(w_0)$.

Experience

e_0: Years of work experience accumulated prior to the most recent interruption.

e_1: Years of work experience accumulated since the last interruption.

e: Total years of work experience ($e = e_0 + e_1$).

TEN: Years of job tenure.

Nonexperience

h_0: Years of nonparticipation prior to the most recent work interruption.

h_1: Duration of the most recent work interruption (in terms of years).

h: Total years of nonparticipation ($h = h_m + h_1$).

"Other" Variables

MAR: Dummy variable = 1 if marriage took place during or immediately before the most recent interruption.

DIV: Dummy variable = 1 if divorce took place during or immediately before the most recent interruption.

BAB: Dummy variable = 1 if a new child was born during or immediately before the most recent interruption.

HLTH: Dummy variable = 1 if a health problem arose during or immediately before the most recent interruption.

MIG: Dummy variable = 1 if migration took place during or immediately before the most recent interruption.

LYOF: Dummy variable = 1 if a layoff occurred during or immediately before the most recent interruption.

UNEM: Dummy variable = 1 if an episode of unemployment occurred during or immediately before the most recent interruption.

RCLL: Dummy variable = 1 if after the most recent interruption the respondent returned to the same job held before that interruption.

APPENDIX TABLE 1
SUMMARY OF STATISTICS: MEANS AND STANDARD DEVIATIONS (IN PARENTHESES)

Sample[a]	I	II	III	IV	V	VI
w_0: ¢ (1967 prices)	—	—	213.1	177.8	—	—
			(98.2)	(97.0)		
w_1: ¢ (1967 prices)	—	—	210.6	150.3	—	—
			(99.9)	(75.9)		
$\ln(w_0)$:			5.24444	5.0129		
			(.592)	(.762)		
$\ln(w_1)$:	5.1922	4.9532	5.2319	4.8653	—	—
	(.5778)	(.6573)	(.551)	(.687)		
$\ln(w_1) - \ln(w_0)$:	—	—	−.0125	−.1475	—	—
			(.431)	(.498)		
$\ln(w_{74})$:	—	—	—	—	5.731	5.703
					(.450)	(.436)
s: (years)	11.83	11.65	11.86	11.55	12.01	12.01
	(2.22)	(2.12)	(2.28)	(2.25)	(2.24)	(2.14)
e_1: (years)	—	—	—	—	9.40	8.83
					(6.10)	(5.82)
e: (years)	11.67	8.25	—	—	18.58	17.66
	(7.03)	(6.14)			(7.22)	(7.01)
h_0: (years)	10.86	12.25	—	—	10.09	11.07
	(6.62)	(6.22)			(7.17)	(6.95)
h_1: (years)	1.30	3.15	.934	3.27	0	0
	(1.86)	(1.59)	(1.71)	(1.61)		

TEN: (years)	3.63	.51	—	—	7.16	6.77
	(5.26)	(1.27)			(6.32)	(6.00)
LYOF: (dummy)	.102	.144	.089	.142	—	—
UNEM: (dummy)	.086	.101	.086	.017	—	—
MAR: (dummy)	.032	.026	.041	.046	—	—
DIV: (dummy)	.071	.062	.070	.054	—	—
BAB: (dummy)	.122	.183	.116	.204	—	—
HLTH: (dummy)	.182	.221	.188	.252	—	—
MIG: (dummy)	.182	.232	.172	.252	—	—
N:	1485	612	1304	373	1015	820

a *Sample I:* Includes all white women married spouse present (at least in part of the survey period) who were either intermittent workers or continuous workers, and, in addition, satisfied the following conditions. Only those intermittent workers are included which have experienced a complete spell of labor force interruption within the survey period (1966–1974) followed by a spell of gainful employment for which reported wage rates are available in the data. Continuous workers are included providing their 1972 wage rates are reported in the data. *Sample II:* is a subsample of I which includes only intermittent workers, and excludes continuous workers. *Sample III:* includes all white women married spouse present (at least in part of the survey period) who were either intermittent workers reporting wages before and after at least one complete spell of labor force interruption within the survey period; or, alternatively, were continuous workers reporting wages in both years, 1971 and 1972. *Sample IV:* is a subsample of III which includes only intermittent workers, and excludes continuous workers. *Sample V:* includes all white women (regardless of marital status) who have reported wages in 1974. *Sample VI:* is a subsample of V which includes only married women husband present.

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Erratum

An error appeared in the Auster and Oaxaca article, "Identification of Supplier Induced Demand in the Health Care Sector," in the Summer 1981 issue. The sentence, beginning on line 16, page 33, should have read: "If an instrumental variables technique (INVR) were used to replace $\ln E$ by its fitted values, the estimate of b_3 would necessarily equal unity."